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Government Green Procurement Spillovers: Evidence from Municipal Building Policies in California[*]

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ABSTRACT

We study how government green procurement policies influence private-sector demand for similar products. Specifically, we measure the impact of municipal policies requiring governments to construct green buildings on private-sector adoption of the US Green Building Council's Leadership in Energy and Environmental Design (LEED) standard. Using matching methods, panel data, and instrumental variables, we find that government procurement rules produce spillover effects that stimulate both private-sector adoption of the LEED standard and investments in green building expertise by local suppliers. These findings suggest that government procurement policies can accelerate the diffusion of new environmental standards that require coordinated complementary investments by various types of private adopter.

JEL Codes: L15, Q58, Q55, O33.

Keywords: Public procurement, green building, quality certification, environmental policy.

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Governments often use their formidable purchasing power to promote environmental policy objectives. The US Environmental Protection Agency and the European Union, for example, have developed environmentally preferable purchasing guidelines for goods ranging from paint, paper, and cleaning supplies to lumber and electricity. Various state and local governments have taken similar steps.[1] These procurement policies often have the stated goals of encouraging cost-reducing innovation among suppliers and spurring private demand for green products (Brander et al. 2003; Marron 2003). The European Union, for example, justifies its environmental procurement policy not only on the basis of leveraging government demand to "create or enlarge markets for environmentally friendly products and services" but also on the basis of stimulating "the use of green standards in private procurement" (Commission of the European Communities 2008: 2). To date, there has been little evidence on whether these targeted government procurement policies produce the intended spillover effects. This paper provides some initial evidence by measuring the impact of municipal green building procurement policies on the private-sector adoption of green building standards.

We examine whether green building requirements that apply *only* to municipal buildings accelerate the use of green building practices by private-sector developers in the same geographic markets, as manifested by more rapid diffusion of the US Green Building Council's Leadership in Energy and Environmental Design (LEED) standard for sustainable building practices. Our results show that the LEED standard diffuses more quickly among private-sector developers in cities that adopt a government green building procurement policy, as compared to a matched sample of non-adopting cities of similar size, demographics, and environmental preferences. We also find that government green procurement policies are associated with the growth of green building input markets, as measured by the number of local architects,

[1] Many authors have discussed government procurement as a policy instrument. For example, see Johnstone (2003), Coggburn and Rahm (2005), Commission of the European Communities (2008), Michelsen and de Boer (2009), and National Association of State Procurement Officials and Responsible Purchasing Network (2010). For examples related to green procurement policies, see Clinton (1998), Commission of the European Communities (2008), Environmental Law Institute (2008), Patrick (2009), Rainwater (2009), National Association of State Procurement Officials (2010), and United Kingdom Office of Government Commerce (2010).

contractors, and other real estate industry professionals who obtain the LEED Accredited Professional designation. Finally, we show that green building procurement policies produce geographic spillovers. In particular, there is more LEED adoption by developers and real estate industry professionals in "neighbor cities"—those bordering a city that has adopted a green building policy—than among these neighboring cities' own set of matched controls.

The paper considers three mechanisms by which municipal green procurement policies could promote diffusion of the LEED standard within the private sector. First, government procurement policies might stimulate local demand for green buildings by raising awareness of buildings' impact on the environment or legitimating a particular standard for measuring green building performance. Second, government procurement policies might lead to lower prices for green building inputs through some combination of increased entry by new suppliers, scale economies, and learning effects. And third, government procurement policies might solve a coordination problem in the market for green buildings. Specifically, if developers are waiting for key suppliers to invest in green building expertise, while those same suppliers are waiting for evidence of ample demand, municipal government procurement policies might jump-start the development of specialized input markets by providing a guaranteed source of demand for LEED-accredited professionals and other suppliers.

While these three mechanisms are not mutually exclusive, our analysis suggests that green building procurement policies promote entry by input suppliers, thereby helping to solve coordination problems associated with joint adoption. In particular, we find no evidence that procurement policies had a larger impact in "greener" cities, which would have supported the theory that procurement policies increase awareness of the LEED standard, and would therefore produce a larger response in markets with greater latent demand for green buildings. We also find that procurement policy impacts are, if anything, somewhat larger in large municipalities. This contradicts the hypothesis that procurement policies cause incumbent green building input suppliers to reduce prices in response to learning, scale economies, or increased competition, since we would expect these effects to be stronger in small markets where competition is weak and suppliers have not reached efficient scale.

This leaves entry by new suppliers and solving coordination problems as two possible explanations for the spillovers we observe. We find that more new suppliers enter (by obtaining LEED accreditation) in markets where there is a local green building procurement policy. While we do not provide a direct test of the coordination hypothesis, the final step in our analysis uses instrumental variables to measure the causal impact of the supply of LEED Accredited Professionals on private developers' LEED adoption rates and vice versa. There can only be coordination failures in LEED adoption if both effects are positive, which we find to be the case.[2] Overall, our findings suggest that government purchasing policies can break deadlocks that emerge when coordinated investments are required to adopt a common standard, thereby stimulating the growth of private markets for the targeted goods and services.

Related literature. This study contributes to four broad literature streams. First, we add to a nascent literature that characterizes how governments are increasingly incorporating environmental criteria into their procurement policies. Much of this work is descriptive. For example, Coggburn and Rahm (2005) and May and Koski (2007) describe the emergence of green building procurement policies within the US federal and state governments. McCrudden (2004) provides an historical context by recounting how governments have used procurement policies to promote a host of social objectives. Michelsen and de Boer (2009) and Sourani and Sohail (2011) identify both the barriers to implementing green building procurement policies and the capabilities that can overcome those barriers. Marron (1997) and Marron (2003) describe the potential impacts of government green procurement policies.

We also contribute to a literature that examines the adoption and impact of green building practices. Kahn and Vaughn (2009) show that LEED certification and Toyota Prius ownership

[2] To estimate the impact of an increase in the number of LEED Accredited Professionals on private developers' LEED adoption rates, we use green building policy adoption in distant cities as an instrument for the number of LEED Accredited Professionals in nearby cities. To show that private developers' LEED adoption rates cause an increase in the supply of LEED Accredited Professionals, we use new construction starts (conditional on city size) to instrument for the level of LEED adoption.

were highly concentrated in wealthy coastal areas. Eichholtz, Kok, and Quigley (2010) provide the first large-scale evidence of private benefits from green building, using building-level data to show that green-certified properties have higher rents and occupancy rates than comparable noncertified properties in the same neighborhood. Kok and Jennen (2012) report similar results. Kok, McGraw, and Quigley (2011) reveal a positive association between the number of LEED Accredited Professionals and the growth rate of LEED certification.

Unlike prior studies of LEED diffusion, our research emphasizes spillovers from public procurement rules to private adoption. Choi (2010) finds greater commercial LEED adoption in cities with municipal policies that provide formal administrative benefits for green building proposals (such as quicker review cycles) or that require commercial buildings to incorporate green features. We provide evidence of spillover effects on private real estate development even when municipal green building procurement policies *do not* provide explicit rules or incentives to encourage private adoption.

Our study also contributes to the literature on quality certification. While this literature typically emphasizes information problems (see Dranove and Jin (2010) for a review), we focus on the role of network effects in the diffusion of a new standard. When the success of a new quality standard depends on many different actors (such as producers, wholesalers, retailers, and customers), certification programs will resemble a multisided platform, with adoption by one group conferring an externality on the others. Farrell and Saloner (1986) model technology adoption in the presence of network effects and coin the term “excess inertia” to describe the familiar chicken-and-egg coordination problem whereby each side waits for the others to adopt. Corts (2010) applies a two-sided platform perspective to study the diffusion of alternative fuels and shows that government procurement of “flex fuel” vehicles that run on both gasoline and ethanol led to increased supply of ethanol at local filling stations. We follow Corts by measuring the impact of government procurement policies on the supply of complements, which in our setting is the number of LEED-accredited professionals. We extend his analysis by measuring the "same-side" externalities (that is, the impact of government procurement policies on private LEED adoption) and by evaluating a broader range of potential mechanisms.[3]

Finally, our examination of the efficacy of government procurement contributes to a growing literature evaluating alternative regulatory approaches such as government voluntary programs (Toffel and Short 2011), voluntary agreements (Delmas and Montes-Sancho 2010), and mandatory information disclosure programs such as restaurant hygiene grade cards (Jin and Leslie 2003).[4] With procurement becoming an increasingly popular policy instrument in Europe (Commission of the European Communities 2008) and in the United States (National Association of State Procurement Officials and Responsible Purchasing Network 2010), our research confirms the promise of this approach, at least in the context of green building.

The rest of the paper is organized as follows: Section I outlines a simple framework for analyzing the impact of green building procurement policies on the private sector and describes the LEED green building standard. Section II describes our data, measures, and empirical methods. Section III describes the empirical results. Section IV offers concluding remarks.

I. Public Procurement and Environmental Standards: Theory and Institutions

A. *Procurement Spillovers in Theory*

Government purchasing guidelines often use price preferences or quantity targets to reward products that meet environmental criteria such as incorporating recycled content, exhibiting pollution levels below regulatory limits, or exceeding voluntary energy-efficiency standards. These policies can significantly boost demand for the targeted products and services through the government's own procurement decisions, especially when the government is a major customer. However, the impact may extend beyond this direct effect, depending on how government purchasing interacts with private-sector demand. Governments often try to design policies that will "influence the behavior of other socio-economic actors by setting the example,

[3] From the literature on multisided platforms, we borrow the "same side" terminology to denote an externality between two groups of users that do not transact with one another but typically use a standard or platform in a similar way (see Rysman 2009).
[4] For a review of this literature, see Doshi, Dowell, and Toffel (2013).

and by sending clear signals to the market-place" (Organisation for Economic Cooperation and Development 2000: 20).

In principle, government procurement policies can influence private-sector purchasing through supply channels, demand channels, or both. Moreover, the private-sector response to a government green purchasing policy might either reinforce or counteract that policy's direct impacts.

Supply channels. On the supply side, government green procurement policies may stimulate private-sector demand for the targeted products and services if increased government purchasing reduces suppliers' average costs; for example, when there are significant scale economies or learning-curve effects in key input markets.[5] When suppliers' fixed costs are large relative to the size of the market, government purchases might also spur entry, leading to more competition and lower prices.

An alternative theory of positive procurement spillovers is that explicit government preference for a particular product or standard will help private market participants overcome excess inertia in the adoption process. By stimulating the supply of goods that meet a particular standard, government demand can provide a focal point for private demand. This theory assumes that private suppliers and customers will not independently adopt a common standard in order to realize the benefits of a more coordinated supply chain, perhaps because of the risk that prior investments in specific standards will be stranded or underutilized.[6] One example of using government policy to overcome this type of coordination failure is the US Department of Agriculture's organic certification program, which was developed partly in response to concerns that farmers and consumers were confused by a proliferation of competing private organic labels and could not coordinate on a common standard (Fetter and Caswell 2002).

[5] For instance, many military technologies require substantial up-front R&D expenditures and rely on the scale economies produced by military procurement programs to reach cost levels that are suitable for civilian application. This theory is closely related to the "induced innovation" hypothesis that procurement preferences lead to increased competition and innovation on the targeted product or service attributes. For example, Siemens (2003) suggests that a preference for the Energy Star label in government computer purchasing led to increased innovation in energy-efficient electronics.
[6] Rochet and Tirole (2006) show that a similar coordination failure is the central assumption in the literature on multisided platforms.

In principle, government procurement policies could also have negative spillovers that stifle private consumption. When supply is inelastic, for example, government procurement might crowd out private purchases of the targeted goods (Marron 1997).[7] Alternatively, if procurement rules define a sharp cutoff between green and brown products, the private supply of environmental goods might become concentrated just above the green-compliance threshold. If some suppliers would have produced greener products in the absence of a sharp cutoff, then environmental procurement rules could actually reduce the supply of green goods, even if they do increase private purchasing of green products.[8]

Demand channels. Government procurement policies might also produce a shift in the private demand curve, as opposed to movement along it. For example, procurement policies could increase the visibility or credibility of a green product (or label) to private consumers, especially when consumers are unable to evaluate claimed environmental benefits on their own. Put differently, procurement policies might unleash latent demand for green goods simply by raising consumer awareness. We expect these information-based demand-side effects to be most salient when the green product or label has minimal market share and little consumer awareness prior to the government's adoption of the procurement policy.

Government procurement rules could also influence private demand by altering the weight that consumers attach to specific policy priorities. For instance, a government could exercise moral suasion, leading private firms and consumers to follow its purchasing guidelines, especially if those parties are already favorably disposed towards the underlying policy goals. On the other hand, public procurement might crowd out private demand if consumers come to perceive that the public sector is already "doing enough" to support those same goals.

[7] While we could find no clear examples of crowding out in green procurement, there is some evidence that the supply of green power is inelastic, so government subsidies for green electricity are primarily spent on marketing and advertising these higher-priced services to end consumers rather than investing in new-generation facilities (Rader 1998).

[8] This seems especially likely when procurement policies are based on voluntary standards developed by firms with strong incentives to preempt more stringent regulation (Lyon and Maxwell 1999; King and Lenox 2000; Reid and Toffel 2009). Interestingly, this suggests that government purchasing policies should sometimes avoid specifying particular private standards, especially when there are questions about the motives of the developers of those standards or about the stringency of the private certification.

Government green building procurement policies. In practice, the importance of any supply- or demand-side channel depends on specific features of that product's market. There are several reasons to expect that, in our analysis, private demand will respond positively to government green building procurement policies. First, government is an especially large customer in the real estate market. According to Marron (2003), 26 percent of all spending on "maintenance and repair construction" comes from federal, state, and local government.[9] Second, builders can realize direct benefits from green investments that produce energy savings or that increase tenants' willingness to pay (Eichholtz, Kok, and Quigley 2010). Third, our analysis covers a period when LEED was just emerging as the dominant standard for green building certification, so government procurement policies could plausibly jump-start key input markets if suppliers were waiting for private developers to commit to a standard. While each of these factors suggests that we should observe a positive correlation between government green building procurement policies and private-sector green building certification, they also suggest that we should be cautious about extrapolating our findings to settings in which the government's share of purchases is small, there are few direct benefits of investment, and standards and technologies are already mature.

B. *LEED Certification and Accreditation*

LEED is a green building certification program developed and administered by the nonprofit US Green Building Council (USGBC). Started in 1998, LEED initially focused on rating the environmental attributes of new construction and has since added rating schemes for commercial and retail interior design, residences, neighborhoods, and building renovation. Federal, state, and local governments have been significant LEED adopters since the program began.

LEED awards points for incorporating specific design elements or meeting environmental

[9] The munitions industry is the only case in which government purchasing represents a larger share of total expenditures.

performance targets in eight categories.[10] More total points qualify projects for increasingly prestigious certification levels: certified, silver, gold, and platinum.

The LEED certification process begins with the developer registering a project with USGBC. Registration “serves as a declaration of intent to certify” the building, provides the developer access to LEED information and tools, and lists the project in the publicly available online LEED project database (Green Building Certification Institute 2011). Once the construction or renovations have been completed and the certification application has been approved, the applicant is sent a plaque (often displayed in the lobby in commercial buildings) and the project is included in the online LEED database of certified projects.

The cost of adopting the building practices necessary to obtain LEED certification varies with the type and scale of the project and with the certification level. Costs can accrue from coordinating the required design elements and from using more expensive materials and technologies. The activities required to obtain LEED points range from relatively cheap (such as installing bike racks) to quite expensive (such as remediating a brownfield site). The administrative costs of LEED certification are small by comparison: roughly $450-600 to register a project with USGBC and a certification fee of $2,000. Estimates of the non-construction-and-materials marginal costs of LEED (“soft costs” that mainly comprise additional design and documentation) range from $0.41 to $0.80 per gross square foot (GSF), or roughly $30,000 for the median project in our sample of LEED buildings.[11]

The benefits of LEED can include increased rents and occupancy rates and reduced operating costs. Several studies have found that LEED-certified buildings charge a three-to-five-percent rent premium and have higher sale prices and occupancy rates (Eichholtz, Kok, and Quigley 2010, 2013; Fuerst and McAllister 2011a, 2011b; Chegut, Eichholtz, and Kok 2013).

[10] The eight LEED categories are: location and planning, sustainable sites, water efficiency, energy and atmosphere, materials and resources, indoor environmental quality, innovation and design, and regional priority.
[11] Estimates of soft costs were obtained from the “LEED Cost Study” commissioned by the US General Services Administration (Contract No. GS–11P–99–MAD–0565, p. 187). Our $30,000 estimate is simply the midpoint of the GSA range ($0.60/GSF) multiplied by 50,000 GSF, which is roughly the median size of a LEED project (the mean project is 216,000 GSF).

Evidence of reduced operating costs is mixed, however, in part because LEED certification emphasizes design elements rather than energy consumption. Nevertheless, several engineering studies do suggest that LEED certification is correlated with increased energy efficiency (Turner and Frankel 2008; Newsham, Mancini, and Birt 2009; Sabapathy et al. 2010).[12]

Given the potential costs and complexity of adopting the LEED standard, the USGBC also created a program to educate and certify real estate industry professionals in its application. To become a LEED Accredited Professional, an individual must pass an exam demonstrating deep knowledge about green buildings in general and the LEED standard in particular.[13] LEED Accredited Professionals come from a variety of occupational categories—including architects, engineers, and project managers—as illustrated in Figure 1.[14]

C. *Empirical Roadmap*

Our analysis of LEED diffusion builds on the idea that the standard resembles a multisided platform that facilitates interactions among real estate developers and suppliers of green building inputs. Thus, our first set of empirical results measures the strength of same-side spillovers in LEED adoption between private developers and government. Specifically, we find a positive relationship between the adoption of government green building procurement policies and the number of LEED-registered private-sector buildings. This relationship could exist for a variety of reasons, including demonstration effects, moral suasion, scale economies, learning effects, anticipation of regulatory changes, or a positive correlation between policies and preferential treatment of green buildings in the municipal permitting process. We conduct several analyses to test these explanations, such as examining whether green building procurement policies have a greater impact in larger cities or in greener cities. We also estimate the policies'

[12] For example, engineering estimates from a study of 121 LEED-certified projects that volunteered data on energy use suggest that these buildings consume 25-30 percent less energy than the national average for comparable projects (Turner and Frankel 2008), though other observers have raised concerns that some LEED-certified buildings do not deliver energy savings (Navarro 2009).

[13] In 2004, it cost roughly $350 to take the LEED Accredited Professional exam. It is hard to assess the opportunity costs of preparation, but one website (www.leeduser.com/) suggests taking a test-prep course and studying an additional 20 hours.

[14] Future research could explore the diffusion and agglomeration of different types of human capital that acquire LEED expertise and how this might differentially affect the diffusion of LEED certified buildings.

impact on private-sector LEED registration rates in neighboring non-adopting cities, where private developers could benefit from geographic spillovers produced by a nearby green building policy. This neighbor city analysis compares LEED registration rates in cities adjacent to green building procurement policy adopters with registration rates in these neighbors' own set of matched control cities. The neighboring city analysis also provides a larger and more representative sample of "treated" municipalities, and helps to address lingering concerns about other omitted variables, such as green preferences in the permitting process, that could be correlated with both policy adoption and private-sector LEED building.

Our second set of empirical results measures the strength of "cross-side" spillovers in LEED adoption between developers and building-industry professionals.[15] For any platform, a larger installed base on one side should generate an increased supply of complements on other sides. We show that government green building procurement policies stimulate investment in green building expertise among local real estate industry professionals (measured as the number of LEED Accredited Professionals). Because the market for real estate professionals often extends beyond city borders, we also examine the impact of green building procurement policies on the supply of LEED Accredited Professionals who work in neighboring cities. This analysis reveals that green building policies increase the supply of green building professionals beyond the policy adopter's city limits. Thus, our results suggest that LEED Accredited Professionals are a key transmission mechanism for geographic spillovers of green procurement policies.

In principle, professional service providers might invest in green building know-how without any government encouragement or formal certification program if they expected such human capital to be rewarded in the marketplace. However, uncertainty about whether and how the market will observe, measure, and reward green building expertise creates a possibility of stranded investment and thus an opportunity for government policies to solve the resulting coordination problem. Moreover, although our cross-side spillover results focus on LEED

[15] In the literature on multisided platforms, a "cross-side effect" is a positive externality between two groups that use a platform to interact with one another. In the literature on network effects, video game players and developers is a leading example.

Accredited Professionals, government green building procurement policies could jump-start many other complementary input markets. For instance, producers and local distributors of building materials might be more likely to carry products that meet LEED criteria after a green building procurement policy is adopted. Viewing the number of LEED Accredited Professionals as a proxy for a host of specialized green inputs helps clarify why developers might be slow to adopt LEED even if they believe there is latent demand for green buildings: The cumulative expense of being a green first-mover could be large, even if contractors and architects constitute a small share of total construction costs.

In our final set of analyses, we switch from measuring the reduced-form impacts of government green procurement policies to measuring the structural links between each side of the LEED platform. In particular, we estimate the causal impact of the number of LEED Accredited Professionals on private-sector LEED registrations by using "distant" green procurement policies as an instrumental variable. The key maintained assumption in this analysis is that municipal green procurement policies in far-away cities increase the supply of LEED Accredited Professionals in nearby markets, but otherwise have nothing to do with the decisions of private developers in a focal market to adopt LEED. To estimate the causal impact of LEED registrations on LEED Accredited Professionals, we use the number of new buildings constructed between 2003 and 2007 (conditional on city size) as an instrument for registrations. We find that both of these structural relationships are positive and somewhat larger than the reduced-form relationships, which supports the theory that government procurement policies promote LEED diffusion by helping real estate developers and building-industry professionals overcome excess inertia in the early stages of adoption.

II. Data and Measures

Our analysis uses data on 735 California cities from 2001 to 2008. We selected California because it is the state with the largest number of municipal green building policies and it also has many cities that had not adopted such policies during our sample period. Our dataset combines information from a variety of sources. We obtained LEED diffusion data from the USGBC, data on nonresidential construction starts from McGraw Hill, and city-level demographic data from the US Census. We hand-collected data on the municipal adoption of green building policies. Summary statistics are presented in Table 1. The unit of analysis is a city (or city-year), which we defined as a Census Place, the geographical unit with available Census demographics and voting-records data that most closely resembles the political unit of a municipality.

LEED registrations. We measure private-sector and public-sector LEED diffusion via LEED registration data obtained from the USGBC. *Annual Private LEED Registrations* is a count of new privately owned nonresidential or multi-unit residential buildings that were registered for LEED certification in a given year;[16] it reflects private-sector developers' intention to use green building practices.[17] This total ranged from 0 to 52 across all the city-years in our sample (which excludes Los Angeles, San Francisco, San Diego, and San Jose)[18] and averaged 1.32 per city-year during our sample period. *Total Private LEED Registrations* is the total (cumulative) number of *Annual Private LEED Registrations* for each city during our sample period of 2001 to 2008. This total ranged from 0 to 87 across all the cities in our sample and averaged 1.64 per city during our sample period. To be clear, it is possible that developers adopt green building practices without obtaining LEED certification. However, given the small marginal costs of certification (conditional on building green) and the evidence that certification is associated with increased rent and occupancy rates, we expect that most qualifying structures actually do seek certification.

[16] We include registrations pertaining to any version of the LEED standard, including those for new construction (LEED-NC), for commercial interiors (LEED-CI), and for a building's core and shell (LEED-CS).
[17] LEED registration is only the first step towards certification. The USGBC encourages projects to register early, since many decisions that will influence certification levels must be taken in the early stages of development. Because the lag from registration to certification can be several years and the LEED standard was diffusing rapidly toward the end of our sample period, a count of certified buildings would have excluded a large number of the projects in our dataset. For the buildings for which we have certification data, the average lag between registration and certification is between two and three years. Anecdotal evidence suggests that few registered buildings fail to certify at some level.
[18] We exclude the four largest cities in California when calculating these summary statistics, since they (a) could not be matched for the analysis below and (b) tend to distort the sample averages due to their extreme size.

LEED Accredited Professionals. Our second outcome measure captures LEED-specific human capital investments by local real estate professionals. *Annual LEED Accredited Professionals* is the number of building industry professionals (such as architects, contractors, and consultants) who passed the USGBC's LEED accreditation exam in a given year. We obtained the city locations of LEED Accredited Professionals from their business addresses maintained in the USGBC directory of LEED Accredited Professionals. *Total LEED Accredited Professionals* is the total (cumulative) number of *Annual LEED Accredited Professionals* during 2001 to 2008; that is, the number of professionals in a city that had become LEED Accredited Professionals during this eight-year period. By 2008, there were between 0 and 416 such professionals in each city in our estimation sample, with an average of 7.5 per city.

Government procurement policies. Our main explanatory variables indicate whether or not a focal city (or a "neighbor city" bordering a focal city) had, by the current calendar year, adopted a municipal green building policy targeting only government buildings. We gathered this policy information by hand, starting from lists compiled by the USGBC and by the Database of State Incentives for Renewables and Efficiency (DSIRE), funded by the US Department of Energy.[19] We identified 155 US cities (40 in California) that had adopted some type of green building ordinance by 2008. (The California green building policy adopter cities in our sample are listed in Appendix Table A-1.) The sample ends in 2008 because of increased regulation of private green building practices—including California's statewide Green Building Policy, which went into effect in August 2009—and because of data availability.

Municipal green building policies vary along several dimensions, including the types of structure affected (by size, owner, and use); whether they cover only new buildings or also renovations; and how they measure environmental performance.[20] We gathered details on each

[19] We acknowledge the excellent research assistance provided by Mark Stout. The DSIRE list of state and local incentives is available at http://www.dsireusa.org/ and the USGBC list can be found at http://www.usgbc.org/PublicPolicy/SearchPublicPolicies.aspx?PageID=1776.

[20] For example, the policy adopted by Irvine, CA in 2005 required new public construction over 5000 square feet to become LEED certified, while the policy adopted by Santa Cruz in the same year had no size threshold and required buildings to achieve a LEED silver certification. While our database is too small to explore the potential for heterogeneous treatment effects based on

policy from city websites and the online library of municipal codes.[21] Our research indicates that 87 percent of all green building polices contained a purchasing rule—that is, a requirement that new public projects adhere to some type of environmental standard—and that 90 percent of these rules specified the LEED standard. Most policies in our sample apply to new construction and do not require buildings rented by municipalities to be certified as green.

We create a time-invariant indicator variable, *Green Policy Adopter*, that equals 1 if a city had adopted a green procurement policy by 2008 and equals 0 otherwise. For cross-sectional models, we create a time-invariant variable, *Exposure to Policy*, to denote the total number of years that had elapsed by 2008 since a city had adopted a policy; it is coded 0 for cities that did not adopt a policy. For panel data models, we create a time-varying indicator variable, *City Adopted Green Policy*, coded 1 starting the year a city adopted a green procurement policy and 0 before that (and always coded 0 for cities that did not adopt such a policy during our sample period). We also create *Years Since City Adopted Green Policy* to count the years since adoption; this, too, was always coded 0 for non-adopting cities. Similarly, for the neighboring city analysis, we create (a) a time-invariant indicator, *Green Policy Adopter Neighbor*, coded 1 for cities that had not adopted a green procurement policy but bordered a city that had done so by 2008, and coded 0 otherwise; (b) a time-varying indictor, *Neighbor Adopted Green Policy*, coded 1 for cities that had not adopted a green procurement policy but bordered a city that had done so by the focal year, and coded 0 otherwise; (c) *Years Since Neighbor Adopted Green Policy* to count the years since a neighbor city adopted a policy and coded 0 in pre-adoption years and for cities that did not neighbor a policy-adopting city;[22] and (d) *Exposure to Neighbor's Policy*, which denotes the number of years that had elapsed by 2008 since a neighbor city adopted a policy and is coded 0 for cities that did not border a policy-adopting city. Four percent of the cities in our estimation

different policy features, future research could attempt such analysis.

[21] Available at www.municode.com.

[22] Seven California municipalities had green building policies that imposed green building mandates on private-sector development in addition to mandates on government buildings. We exclude these seven cities from our analysis of *Green Policy Adopter* cities, but treat them as green building procurement policy adopters when analyzing *Green Policy Adopter Neighbor* cities.

sample had adopted a municipal green building policy by 2008 and 15 percent of the cities in our sample are green policy adopter neighbors.[23]

Construction activity. To control for variation in the underlying rate of new building activity, we purchased quarterly data on new building starts from McGraw Hill's Dodge Construction Reports between 2003 and 2007.[24] The control variable *Annual New Buildings* is the annual number of nonresidential construction starts in each city. For periods for which we do not have data on new construction starts, we extrapolate based on the nearest six preceding/following quarters of new construction starts. *Total New Buildings* is the cumulative count of nonresidential construction starts between 2003 and 2007. The mean number of *Total New Buildings* for a city in our estimation sample was 26.21. Since the *Total New Buildings* variable is highly skewed and strongly correlated with population ($\rho = 0.88$), we also calculated the number of *Total New Buildings per Capita* (measured in buildings per 10,000 residents), which has a mean of 12.06 in our sample.

Demographics. For each city in the analysis, we obtained *Population* (measured in units of 10,000), *Income* (median household income in $10,000s), and *College* (the share of adults with some college education) at the Census-Place level from the 2000 US Census. To create a proxy for white color employment, we aggregated the employment of all establishments in the finance, insurance, and real estate (FIRE) industries in 2001 from the National Establishment Time Series (NETS) database, a compendium of Dun & Bradstreet data, to create *FIRE employment*.

Environmental preferences. We collected several measures of the prevailing preference for environmental sustainability in each city. First, we gathered data from the University of California's Statewide Database (http://swdb.berkeley.edu/) and calculated *Green Ballot Share* as the proportion of each city's citizens in favor of various statewide environmental ballot

[23] While our matching procedure (described below) excludes Los Angeles, San Francisco, San Diego, and San Jose from the analysis of procurement policy adopters, each of these cities did in fact adopt a green building procurement policy and we treat them as policy adopters in the neighbor city analysis.
[24] We could only afford to procure five years of building-level construction starts data for California.

initiatives addressing environmental quality during 1996-2000 (Kahn 2002; Wu and Cutter 2011). These ballot initiatives received support from an average of 61 percent of each city's citizenry.

Second, we obtained data on green purchasing behaviors by calculating the proportion of vehicles registered in 2008 that were Toyota Priuses, based on ZIP-code-level vehicle registration data from RL Polk (Kahn and Vaughn 2009; Kahn 2011). We aggregated these registration data to the city level to reflect the Prius market share in each city, creating the variable *Prius Share*, which has a mean of 0.54 percent.[25] From the Environment California Research and Policy Center, we collected data on the number of residential, commercial, and government solar installations in each city completed by 2006, creating the variable *Solar Projects*.[26] We created *Alternative Fuel Stations*, the number of alternative-fuel filling stations in each city in 2003, from the US Department of Energy Alternative Fuels Data Center.[27]

Finally, using data from the League of Conservation Voters (LCV), we calculated the proportion of pro-environment votes on environment-related bills cast by each city's delegates to the State of California's Senate and Assembly. These variables, *LCV Senate Score* and *LCV Assembly Score,* range from 0 (for cities whose delegates voted against all environment-related bills) to 100 (for cities whose delegates voted in favor of all such bills), with an average near 50 for both the Assembly and the Senate across all cities in our estimation sample.[28]

III. Analysis and Results

A. *Matching and Balance*

To estimate the causal impact of government green building procurement policies on private-sector LEED registrations and LEED Accredited Professionals, we construct a matched

[25] The highest Prius registration rate is 3.74 percent in Portola Valley (just west of Palo Alto).
[26] The raw data on solar projects is found in the public report available at http://www.environmentcalifornia.org/sites/environment/files/reports/California's%20Solar%20Cities%202012%20-%20Final.pdf.
[27] http://www.afdc.energy.gov/.
[28] We use scores from 2001 in our cross-sectional models and annual values in our panel models.

control sample using the coarsened exact matching (CEM) procedure developed by Iacus, King, and Porro (2011, 2012). This method assumes selection on observables. That is, after matching and reweighting the data to account for the joint distribution of observed exogenous variables, we assume that adoption of a green building procurement policy by a city (or its neighbor) is uncorrelated with all other factors that influence private-sector LEED adoption. Intuitively, CEM is a method of preprocessing a dataset before running a weighted least-squares regression and it resembles propensity score methods in its use of matching, sampling weights, and balancing tests.[29]

To implement CEM, one begins by selecting a set of variables on which to match, "coarsening" (discretizing) any continuous variables in the set, and creating a group of "cells" corresponding to all possible combinations of values of the coarsened variables.[30] The set of matching variables and cut points are chosen by the analyst to balance a tradeoff between bias and variance. Adding variables and cut points leads to closer matches in the values of the continuous variables, but also discards more data. The next step in the CEM process is to discard observations from any cell that does not contain both treated and control observations. Finally, a weight of 1 is assigned to each treated unit and a weight of T_i/C_i is assigned to each control observation in cell *i* (where T_i and C_i are the number of treatment and control observations, respectively). Weighted least-squares estimation then yields an estimate of the treatment effect for treated cities remaining in the estimation sample.

Iacus, King, and Porro (2012) describe several advantages of CEM over the propensity score and other matching techniques. Unlike conventional regression control methods, CEM does not extrapolate counterfactual outcomes to regions of the parameter space where there are no data on controls. Because CEM is nonparametric, there is no possibility that a misspecified model of selection will produce greater imbalance in variables that are omitted from the

[29] Step-by-step guidance on implementing coarsened exact matching is provided in Blackwell et al. (2009).
[30] Thus, if there are K matching variables and each (coarsened) variable has L_k possible values, the number of unique cells (prior to discarding any cells that contain no matches) will be $L_1 \times L_2 \times L_3 \times \ldots \times L_K$.

matching procedure, which can happen with the propensity score. Moreover, CEM ensures that the reweighted control sample matches *all* of the sample moments of the treated sample, not just the means.[31] We chose to use CEM because it appears more robust to misspecification of the selection process and because we find the nonparametric exact matching process more intuitive than an iterative search for an appropriate specification of the propensity score. However, because both CEM and the propensity score rely on the same fundamental assumption that selection into treatment is exogenous conditional on observables, the two methods should produce similar results when that assumption is correct.

We use CEM to construct two matched samples: one consisting of green policy adopters and their quasi-control group and another consisting of green policy adopter neighbors and their quasi-control group. In both cases, our goal is to achieve balance—statistically indistinguishable distributions between the treatments and controls—across a set of exogenous covariates that might lead to policy adoption, including environmental preferences, market size, and other city-level demographics.

To implement CEM, we begin by coarsening *Population* to create 10 strata.[32] This large number of strata (relative to the overall size of the dataset) results in a very close match on the size distribution, but leads to a curse of dimensionality (that is, very small samples) if we include many additional variables in the matching procedure. Therefore, when we match policy adopter cities, we match on *Population* and *Prius Share*, for which we assign cut points at the 25th, 50th, 75th, 90th, and 95th percentiles.[33] Matching on these two variables yields a matched sample of 26 policy adopting cities and 180 controls. Because the green policy neighbors sample is somewhat larger, we also match on *Income*, *Green Ballot Share*, and *LCV Senate Score,* though we use a very coarse match for these additional variables in order to prevent a substantial drop in

[31] This property of CEM proved important in our application, where the distribution of city size is highly skewed.

[32] We set cut points at 10, 50, 70, 100, 120, 150, 250, 300, 350, and 470 thousand inhabitants and omit cities above the top threshold because there are no suitable controls.

[33] In terms of actual registration rates, the corresponding values are 0.5, 0.8, 1.0, 1.5, and 2.7 percent of all registered vehicles.

sample size.[34] We omit green policy adopter cities as potential controls for the sample of green policy adopter neighbors. The matching process removes 31 green policy neighbor cities and 324 potential controls, resulting in a matched sample of 80 green policy adopter neighbor cities and 291 matched control cities.

Table 2 illustrates how CEM dramatically improves the balance in the means of exogenous covariates across the treatment and control samples. Each row in the table reports means for the treatment and control cities in a particular sample and a t-statistic from regressing each covariate on the treatment dummy (*Green Policy Adopter* or *Green Policy Adopter Neighbor*). Panel A of Table 2 compares all cities that adopt a green building policy, excluding the four largest, to the full set of potential controls (that is, to all other cities in California) using unweighted OLS regressions.[35] We find that cities adopting a green building policy are larger, greener, wealthier, and better educated than the potential controls. There is a statistically significant difference in the means of each variable except for the per-capita measure of new construction activity.

Panel B of Table 2 compares CEM-weighted means for the matched sample of green policy adopters and their controls. These results can be viewed as a falsification test for our maintained assumption: If the treatment remains correlated with observables, we might be more skeptical of the assumption that it is uncorrelated with unobserved drivers of private LEED adoption. Note that matching on *Population* and *Prius Share* excludes three cities (Oakland, Berkeley, and Ventura) from the treatment group, reducing it to just 26 green building procurement policy adopters. Since we used the distributions of *Population* and *Prius Share* to create the match, by construction we should observe no difference in the means of these variables across treatment and control cities. In fact, Panel B of Table 2 shows that matching on

[34] For the neighbor-city matching, we leave the *Population* cut points unchanged. We continue to use the 25th, 50th, 75th, 90th, and 95th percentiles of *Prius Share*, which correspond to registration rates of 0.26, 0.56, 1.21, 1.78, and 2.36 percent of all vehicles. Finally, we set cut points at the 25th and 75th percentiles of *Income* ($44 and $70 thousand) and at the medians of *Green Ballot Share* (67 percent approval) and *LCV Senate Score* (44 points).

[35] Each of the four largest cities in California (Los Angeles, San Diego, San Jose, and San Francisco) has adopted a green building procurement policy. Including these cities in the analysis leads to a dramatic increase in imbalance and a similarly large increase in the results presented below.

just these two variables eliminates any statistically significant differences in the means of all observables. In particular, CEM produces balance for alternative measures of green preferences (*Alternative Fuel Stations* and *Solar Projects*) and demographic characteristics (*College*, *Income*, and log*(FIRE Employment*)) that were not used to construct the match.

Green policy adopter neighboring cities, too, are larger, greener, wealthier, and better educated than all non-adopting cities (that is, all of their potential controls). Indeed, the raw means of all exogenous covariates in Table 2 were statistically significantly different between the two groups (not shown). CEM matching and reweighting removed significant differences in the means of all of these variables, as indicated in Panel C of Table 2. Given more data, we might consider adding a number of additional variables to the matching process, particularly for the green policy adopter sample. However, the results in Table 2 suggest that we have removed much of the potential bias, and we do not wish to further increase the variance of our estimates by excluding more observations.

B. *Cross-sectional Analysis*

We begin our empirical analysis with a cross-sectional comparison of cumulative LEED registrations between the matched green policy adopter cities and their control cities. Specifically, we estimate the following linear regression:

(1) $$Y_i = \alpha + \beta \cdot Exposure_i + \boldsymbol{\gamma} \cdot X_i + \varepsilon_i,$$

where Y_i is *Total Private LEED Registrations* in city *i* as of 2008. $Exposure_i$ represents *Exposure to Policy,* the number of years that had elapsed between a city having adopted its green policy and 2008, the final year of our sample. X_i represents a set of controls for factors potentially associated with LEED adoption: environmental preferences (*Prius Share*, *Green Ballot Share*, *LCV Senate Score*, and *LCV Assembly Score*), market size and economic growth (*Population*, *Total New Buildings*), educational attainment (*College*), and wealth (*Income*).[36] We are

[36] Adding controls is not necessary for causal inference, but should increase the precision of our estimates. However, it is worth noting that because of the matching procedure, we do not use the control variables to extrapolate potential outcomes to regions of

interested in the coefficient β, which measures the difference in the average annual LEED registration rate between *Green Policy Adopter* cities and their matched controls. Robust standard errors are clustered by county to account for the possibility of spatial correlation of unknown form. We estimate these models with CEM-weighted OLS regression.[37]

Results are presented in Table 3.[38] Our estimates of the spillover effects of government procurement on private-sector demand are presented in Column 1. We find a statistically significant increase of 2.1 private-sector LEED registrations per year in cities with a green building policy, relative to their matched controls. This represents a 30-percent increase in LEED adoption beyond the weighted mean of 7.4 *Total Private LEED Registrations*.

To estimate the impact of procurement policies on *Total Private LEED Registrations* in adjacent cities, we estimate a model akin to Equation 1, except that we replace *Exposure to Policy* with *Exposure to Neighbor's Policy* and use the matched sample of *Green Policy Adopter Neighbor* cities and their matched controls.[39] This model yields a statistically significant increase of 0.15 private LEED registrations per year among neighbors relative to their matched controls (Table 3, Column 2). When normalized by the weighted mean baseline *Total Private LEED Registration* rate of 1.3 buildings per year, this translates to a marginal effect of 12 percent. These estimates suggest that green building procurement policies produce geographic spillovers that influence private-sector LEED adoption in neighboring cities. The results also suggest that

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the parameter space where there are very few treated or untreated units.

[37] As stressed in Angrist and Pischke (2009), OLS provides the best linear approximation to the conditional expectation function, even though Y_i is a count variable. Estimating a model with an exponential conditional expectation function (i.e., Poisson with a robust covariance matrix) produces similar results.

[38] As a preliminary step, we estimate this model on *Total Government LEED Registrations* to verify that municipal government green procurement policies actually lead to an increase in government LEED procurement. To construct this variable, we first create *Annual Government LEED Registrations* as the count of new nonresidential structures that are owned by a local government and that were registered for LEED certification each year. *Total Government LEED Registrations* is each city's total number of *Annual Government LEED Registrations* from 2001 to 2008. During that time, the cities in our sample registered between 0 and 12 new government buildings, with an average of 0.3 LEED-registered buildings per city. Regressing *Total Government LEED Registrations* on *Exposure to Policy* indicates that government green procurement policies—as intended—spur greater municipal green building. We find an average of 0.56 more government LEED registrations per year in cities once they have adopted a green building procurement policy, a statistically significant difference compared to their control cities (see Table A-2, Column 1).

[39] As a preliminary step, we estimate this model on the number of government LEED registrations. We find an average 23% annual increase in neighboring cities that do not themselves adopt a green building procurement policy, compared to the government LEED registration growth rates in these neighboring cities' matched controls (calculated as β=0.06 divided by the CEM-weighted mean outcome of 0.26 in this matched sample; see Column 2 of Table A-2).

the estimates in Column 1 for *Green Policy Adopter* cities are neither an artifact of preferential treatment for green buildings by local zoning or permitting officials, nor an artifact of policy-adopting cities' preference for green buildings in their own rental market.

To examine the impact of government procurement policies on green building input markets, we reestimate Equation 1, except that Y_i becomes *Total LEED Accredited Professionals* in city *i* as of 2008. Column 3 shows an annual increase of 9.2 LEED Accredited Professionals in green policy adopting cities relative to those cities' matched controls. This is an increase of 22 percent over the weighted sample mean and is statistically significant at the 10-percent level.

Since green building factor markets almost certainly extend beyond the borders of any particular municipality, we also estimate the same model using the matched sample of *Green Policy Adopter Neighbor* cities and their controls. Column 4 in Table 3 presents estimates of the impact of being a green policy neighbor on the number of LEED Accredited Professionals. We find a statistically significant increase of 0.7 LEED Accredited Professionals per year, or roughly 10 percent of the weighted sample mean. This suggests that the market for architects, contractors, consultants, and others with green building capabilities is regional, with spillover from policy adopters to neighboring cities. In fact, if we aggregate the outcome variable used in Column 2 to examine the impact of green building procurement policies on the number of LEED Accredited Professionals in all surrounding municipalities, we find large and statistically significant results. However, we focus on city-level outcomes because it ensures a better match between the treated and control cities, which by construction of the matching process have a similar size distribution.

The results in Table 3 are robust to a variety of changes in model specification. The estimated impact of green procurement policies increases if we use the unmatched sample (results not reported), but changes very little if we drop the CEM weights from the OLS models (Panel A of Appendix Table A-3) or omit the control variables (not reported). Estimating the models via a CEM-weighted Poisson regression with robust standard errors, after taking logs of the explanatory variables, yields the same general insights as our primary model with somewhat more statistical precision (Panel B of Table A-3). Finally, we obtain very similar estimates if we expand the data to an eight-year balanced panel of cities and estimate pooled cross-sectional regressions (with or without CEM weights) of *Annual Private LEED Registrations* or *Annual LEED Accredited Professionals* on the policy-adoption indicator variables *City Adopted Green Building Policy* and *Neighbor Adopted Green Building Policy* (unreported).

C. *Panel Data Analysis*

We now exploit the panel nature of our policy-adoption and outcome measures to estimate models that compare LEED diffusion in treatment and control cities before and after the adoption of a green procurement policy. Specifically, we estimate the following two-way fixed-effects model over the years 2001 through 2008:

(2) $$Y_{it} = \alpha_i + \lambda_t + \beta_1 \, \textit{Years-since-adoption}_{it} + \boldsymbol{\gamma} \cdot X_{it} + \varepsilon_{it},$$

where Y_{it} is either *Annual Private LEED Registrations* or *Annual LEED Accredited Professionals* in city i in year t, α_i is a fixed effect that absorbs all observed and unobserved time-invariant city characteristics, λ_t is a set of year dummies, and X_{it} measures annual nonresidential construction starts in city i in year t. When we analyze the focal cities and their controls, *Years-since-adoption*$_{it}$ represents *Years Since City Adopted Green Policy.* Similarly, when we analyze the neighbor cities and their controls, *Years-since-adoption*$_{it}$ represents *Years Since Neighbor Adopted Green Policy.* The coefficient β_1 measures any trend-change in the rate of LEED diffusion following the adoption of a green building procurement policy by policy adopter cities or their neighbors. This specification is slightly different from a standard difference-in-differences regression, which would replace *Years-since-adoption*$_{\mathrm{it}}$ with the indicator variable *Adoption*$_{it}$. However, regressions that allow both a step-change and a trend-change in the LEED adoption rate typically find the coefficient on *Adoption*$_{it}$ to be statistically insignificant, so we constrain it to equal zero.[40] We estimate Model 2 by CEM-weighted OLS regression and continue to cluster standard errors at the county level.

The results of our panel data models are reported in Table 4.We find a substantial but statistically insignificant positive trend-change in *Annual Private LEED Registrations* among green policy adopter cities compared to their matched controls (Column 1). For green policy adopter neighbors, we find a positive and statistically significant trend-change in *Annual Private LEED Registrations* compared to their matched controls (Column 2). We find a similar pattern of results for *Annual LEED Accredited Professionals*: positive trend-changes that are statistically insignificant among focal cities (Columns 3) but statistically significant among our larger matched set of neighboring cities (Column 4).

The estimates in Table 4 are initially somewhat smaller than those in Table 3, but suggest that the gap in LEED adoption between cities affected by a green building procurement policy and their matched controls increases over time. For example, Column 4 in Table 3 suggests that *Green-Policy Adopter Neighbor* cities generate 0.71 more new LEED Accredited Professionals per year than their matched controls following policy adoption, whereas the results of the more flexible specification reported in Column 4 in Table 4 suggest that the difference is 0.24 additional LEED Accredited Professionals in the year the policy is adopted, 0.48 in the second year, 0.72 in the third year, and so on.

These results in Table 4 are robust to several alternative model specifications (reported in Appendix Table A-4). Dropping CEM weights yields similar point estimates, but all of the coefficients become statistically significant (Panel A in Table A-4). Estimating a CEM-weighted model with the outcome in logs indicates, in both the focal city and neighboring city analyses, a positive trend-change that is significant at the 10-percent level for three out of the four cases (Panel B). Finally, we estimate a first-differenced model, which should alleviate any concern about the strict exogeneity assumption associated with conditional fixed effects, and find

[40] The only exception to this statement relates to Model 2 in Table 4, where the more flexible specification finds a *negative* and significant step-change, leading us to report a somewhat smaller (i.e., less positive) trend-change in Table 4.

somewhat smaller positive trend-changes for both private LEED registrations and LEED Accredited Professionals that remain statistically significant for neighbor cities (Panel C).

Returning to Table 4, the bottom two rows report results of F-tests of the null hypothesis that there is no difference in the trends of the outcome variable between treatment and control cities prior to the adoption of the green building procurement policy. To implement this test, we drop from the estimation sample all observations where *Years-since-adoption*$_{it}$ is greater than zero, add a new set of indicator variables coded 1 *t* years before a city *i* adopts a policy (where *t* equals 1 through 4) and otherwise coded 0, and report an F-test for the joint significance of these pre-policy indicators. The F-tests show that there was no difference in pre-policy LEED Registration trends, but that real estate professionals in green policy adopter neighbor cities were becoming LEED-accredited at a higher rate than that of their matched controls before the policies went into effect.[41] This pre-policy trend in LEED accreditation might reflect the fact that there is typically some public discussion prior to the adoption of a green building procurement policy. Indeed, one interpretation of the results in Tables 3 and 4 is that municipal green building policies help solve a coordination problem between developers and green building input suppliers by providing a highly visible source of demand. The next subsection considers some of the alternative interpretations and mechanisms described above.

D. *Moderating Effects of Awareness and Market Size*

One potential explanation for the findings in Tables 3 and 4 is that developers, consumers, and input suppliers were unaware of LEED certification before their municipal government adopted a green building policy. If a combination of latent demand and increased awareness were driving our results, we would expect to see public green procurement policies having a larger impact in cities where there is a higher demand for other green amenities, since a greater local willingness-to-pay for environmental amenities would raise the private benefits of

[41] Using an alternative hazard specification, we find no significant influence of either cumulative LEED registrations or cumulative LEED Accredited Professionals on the adoption of a government green building procurement policy (results available upon request).

adopting green building practices in these locales (for example, through increased rents and occupancy). We explore this idea by estimating cross-sectional OLS models that predict *Total Private LEED Registrations* based on interactions of *Exposure to Policy* or *Exposure to Neighbor's Policy* with a city's average preference for environmental amenities (*Prius Share* and *Green Ballot Share*) or its demographics (*College*). In each regression, we demean the continuous variable in the interaction, so the main effect of policy adoption can be interpreted as an average treatment effect on the treated.

The results in Table 5 show that *Total Private LEED Registrations* does not exhibit a stronger response to a public green procurement policy in cities with greater *Prius Share*, *College*, or *Green Ballot Share*, as compared to cities with lower values of these proxies for environmental preference. This finding holds both for green policy adopter cities and their matched controls and for green policy adopter neighbors and their matched controls. [42] As one might expect, the main effects of *Prius Share* and *College* are positive and statistically significant. While the *Prius Share* interactions are imprecise, the interactions with *College* and *Green Ballot Share* are essentially zero. *LCV Assembly Score* and *LCV Senate Score* also yielded precisely estimated zeroes on the interaction term (unreported).[43] Alternative specifications, in which we replace the continuous moderators with dummy variables indicating whether or not each moderator was above the sample mean, yield the same general pattern of results (see Table A-5). Overall, the estimates in Table 5 suggest that public green procurement policies are not stimulating latent demand by making consumers, private developers, and suppliers more aware that the LEED standard exists.

[42] We considered a number of alternative specifications and found qualitatively similar results for unweighted OLS regressions with and without controls and for count data specifications (i.e., robust Poisson). While no models indicate that there were more *Total Private LEED Registrations* in policy adopting cities with more pro-environmental voting records, we did find that the treatment effect was larger for policy adopters with a high rate of Prius ownership in models where the outcome variable was the number of *Total Private LEED Registrations per new buildings* constructed between 2003 and 2007.

[43] We considered a number of alternative specifications and found qualitatively similar results for unweighted OLS regressions with and without controls and for count data specifications (i.e., robust Poisson). While no models suggest that there were more *LEED Registrations* in treated cities with a more pro-environmental voting record, we did find that the treatment effect was larger for policy adopters with a high rate of Prius ownership in models for which the outcome variable was the *number of LEED Registrations per New Building* constructed between 2003 and 2007.

Another potential explanation for the results in Tables 3 and 4 is that government procurement rules lead to lower green building input prices, thereby stimulating private-sector adoption. In this scenario, incumbent suppliers reduce their prices either because of a decline in average costs (for example, through scale or learning effects) or because more competitors have entered the relevant factor markets (Bresnahan and Reiss 1991). One testable implication of this hypothesis is that the impact of municipal procurement policies should decline with city size.[44] Intuitively, private demand for LEED buildings is more likely to cover a supplier's entry costs in large markets, leading to robust competition among suppliers operating at efficient scale. Conversely, in large cities with competitive factor markets, the increased demand from a municipal procurement policy will have little or no impact on suppliers' average costs or the prices faced by developers.

We examine whether city size moderates the impact of municipal green procurement policies by estimating cross-sectional OLS models of *Total LEED Accredited Professionals* on three measures of market size (log *City Population*, log *County Population* and log *Total New Buildings*) interacted with each of our two treatment dummies (*Green Policy Adopter* and *Green Policy Adopter Neighbor*). The results of these six models, reported in Appendix Table A-6, suggest that there is little relationship between market size and the impact of green building procurement policy adoption. Specifically, the interaction term was statistically insignificant in five of the six models and was positive and significant in the specification in which we interacted policy adoption with log(*Total New Buildings*) for policy adopter cities.[45] Overall, the estimates suggest that government green building procurement policies do not stimulate private demand by causing incumbent suppliers to lower their prices.

[44] One exception is when procurement policies are insufficient to stimulate supplier entry in the smallest markets, leading to a non-monotonic prediction that policy effects will be greatest at some intermediate market size (below the level at which private demand for green buildings is sufficient to generate robust competition among suppliers).

[45] Once again, the results are qualitatively robust to omitting the CEM weights, using a Poisson specification, or dropping the control variables; that is, no specification suggests that the treatment effect declines with city size.

E. *Indirect Network Effects*

The analyses thus far have estimated the impact of government green procurement policy adoption on private-sector green building activity (registrations) and on LEED-specific investments in human capital for green building (accredited professionals). We find no evidence that these results are driven by increased awareness of LEED in policy adopter cities or by scale economies and increased competition among incumbent suppliers. Another possible explanation for our main findings is that green building procurement policies break a deadlock among specialized input suppliers such as real estate professionals, who are reluctant to invest in LEED without evidence of demand for green buildings, and developers, who are reluctant to embark on building green until local factor markets have matured. Our final empirical analyses estimate the indirect network effects at the heart of this story; that is, the causal impacts of LEED Accredited Professionals on LEED registrations and vice versa. Our theory that procurement policies help local markets overcome excess inertia implies a positive feedback loop, which requires both of these structural parameters to be positive.

Instrumental variable models. We use instrumental variables to estimate these indirect network effect parameters. To identify the impact of *Total LEED Accredited Professionals* on *Total Private LEED Registrations*, we require an instrument that is correlated with the supply of LEED Accredited Professionals but uncorrelated with unobserved drivers of private LEED registration. We propose to use government green procurement policy adoption in "distant" cities as our instrument. Specifically, we use the log of the number of green policy adopter cities between 25 and 50 miles from the center of the focal city to instrument for the number of LEED Accredited Professionals in all cities within 25 miles of that focal city. This instrument is motivated by the assumption that markets served by building industry professionals are more dispersed than both the drivers of municipal procurement policy and the direct impact of green building procurement policies. Put differently, we assume that green building procurement policies in cities that are 25 to 50 miles away have no impact on developers of private buildings other than through the supply of LEED Accredited Professionals.[46]

To isolate the reverse relationship—the impact of the number of private LEED registrations on the number of LEED Accredited Professionals—we require an instrumental variable that is correlated with the number of LEED registrations but uncorrelated with unobserved drivers of local real estate professionals' decisions to seek accreditation. Building on the instrumental variables strategy used in Corts (2010), we use *Total New Buildings* (conditional on city population, which we control for in the regression) as an instrument for *Total Private LEED Registrations*. Intuitively, as the number of new building starts increases, so does the probability of having one or more LEED-registered projects that could induce real estate professionals to seek LEED accreditation. Because we condition on *Population*, the key assumption underlying the validity of our instrument is that variation in the *intensity* of development (that is, the number of new buildings per capita within each city) between 2003 and 2007 will affect the number of private LEED registrations (for example, because of competition among developers) without otherwise altering the incentive for real estate professionals to seek LEED accreditation. Because the number of new buildings is clearly exogenous to an individual real estate professional's decision to seek LEED accreditation, the main concern with this instrument is that omitted variables might be correlated with both building activity and LEED accreditation rates. We therefore continue to control for *Prius Share* and *Green Ballot Share* to account for a city's green taste. All models also control for *Population*, *Income*, and *College*.

We estimate our IV regressions on all cities with more than 20,000 residents that did not adopt a green building procurement policy. We exclude policy adopter cities because variation in neighbor city Accredited Professionals is clearly not exogenous for those cities. We exclude cities with fewer than 20,000 residents for comparability to our prior results, where the matching process excluded most small and/or rural cities.

Instrumental variable results. Columns 1 and 2 of Table 6 report cross-sectional OLS and

[46] We used project-level data on construction starts from McGraw Hill to find the median distance between a building site and the office address of its architect or general contractor at 28 miles. The 75th percentile of this distribution is roughly 75 miles.

IV estimates of the impact of *Total LEED Accredited Professionals* on *Total Private LEED Registrations*.[47] Column 1 reports OLS estimates of the correlation between the number of LEED Accredited Professionals in the cities within 25 miles of a focal city and the number of LEED registrations in that focal city. This correlation suggests an increase of 0.29 private LEED registrations per log-point increase in the number of LEED Accredited Professionals in the surrounding cities. Column 2 presents our IV estimates that use distant policy adoption as an instrument for the number of nearby LEED Accredited Professionals. There is a strong first-stage correlation between distant cities' green building policies and the number of LEED Accredited Professionals in municipalities surrounding the focal city. In particular, an F-test clearly rejects the hypothesis that distant green building policies are uncorrelated with the supply of local LEED Accredited Professionals, with an F-statistic (44.5) well above the F=10 rule-of-thumb for diagnosing weak instruments (Staiger and Stock 1997). The second-stage IV estimates indicate that a one-log-point increase in the number of nearby *LEED Accredited Professionals* generates an additional 0.56 *Total Private LEED Registrations* as of 2008.[48] This IV coefficient is almost twice as large as the corresponding OLS estimate and the second-to-last row in Table 6 indicates that we can reject (at the 10-percent-significance level) the hypothesis that the IV and OLS estimates are equal.

Columns 3 and 4 of Table 6 estimate the impact of *Total Private LEED Registrations* on *Total LEED Accredited Professionals*. OLS results are presented in Column 3 as a baseline. In Column 4, we use *Total New Buildings* as an instrument for *Total Private LEED Registrations*. For this model, we also find a strong first-stage relationship between the instrument and the endogenous regressor, with a first-stage F-statistic of 17.7 once again suggesting that we do not have a weak-instrument problem.[49] The IV estimates in Column 4 indicate that each private

[47] Instrumental variable results for the full sample are reported in Appendix Table A-7.
[48] In an unreported log-log specification, the IV coefficient implies an elasticity of 0.1 and remains statistically significant.
[49] The F-statistics reported in Table 6 are based on cluster-robust standard errors with unspecified county-level correlation. An alternative specification that assumes homoskedastic second-stage residuals produces even larger first-stage F-statistics (F=104.1 and 34.5 for Columns 2 and 4, respectively). In this alternative specification, both second-stage estimates remain significant at the 5-percent level, though the standard error on log*(Total LEED Accredited Professionals within 25 Miles)* in Column 2

LEED registration produces three additional LEED Accredited Professionals in the same city.[50] While the IV estimate in Column 4 is about 42 percent larger than the corresponding OLS coefficient, it is also less precise, and we cannot reject the null hypothesis that the OLS and IV estimates are equal.

The instrumental variable results in Table 6 provide evidence of two positive causal relationships operating simultaneously: (1) An increase in the supply of LEED Accredited Professionals causes an increase in the number of private LEED registrations and (2) an increase in the number of private LEED registrations causes an increase in the number of LEED Accredited Professionals. These indirect network effects are a necessary (though not sufficient) condition for the existence of a chicken-and-egg dilemma in the adoption of a new quality standard.[51] More generally, by showing how distant green procurement policies can influence private-sector developers through the supply of nearby LEED Accredited Professionals, these results point to the importance of supply-side spillovers in the diffusion of LEED.

IV. Discussion and Conclusions

This paper provides evidence that public procurement policies can influence private-sector purchasing decisions in a way that reinforces underlying policy goals. In particular, we show that local-government green-building procurement rules produce spillover effects that stimulate both private-sector adoption of the LEED standard by developers and investments in green building expertise by local suppliers. These effects are analogous to the same-side and cross-side externalities emphasized in the industrial organization literature on technology platforms. Moreover, in our context there is a geographic component to these spillover effects, as the development of regional input markets stimulates private-sector developers in neighboring cities to adopt LEED at a greater rate, even when those neighbor cities do not have their own

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increases to 0.28 while the standard error on *Total Private LEED Registrations* in Column 4 drops to 0.67.

[50] In an unreported log-log specification, the IV coefficient implies an elasticity of 1.24 and remains statistically significant.

[51] Future work using a larger sample of cities might estimate a structural model that explicitly accounts for the possibility of multiple equilibria in the adoption process.

green building policies. To our knowledge, this the first study to examine whether or not government procurement acts as a focal adopter that tips the market towards a particular standard or certification scheme, despite this often being a primary stated objective of socially motivated government procurement policies such as “buy green” initiatives.

Another contribution of our study is to suggest a parallel between standards for environmental performance and the interoperability standards studied in the industrial organization literature on platforms and compatibility. In particular, we find evidence of an excess inertia or chicken-and-egg problem—a type of coordination failure typically associated with hardware-software platforms—in the diffusion of a new environmental standard, and we show how this problem might be overcome if local governments step in as lead users. The industrial organization literature describes a variety of alternative ways that a platform manager could resolve this type of dilemma; for example, by offering aggressive pricing for a particular set of lead users or by supplying its own complements. Future research might focus on cases where these alternative strategies were used to promote adoption of different types of environmental standards. More generally, understanding how insights into platform creation and governance could apply to industry-led efforts to develop environmental standards strikes us as a promising research agenda.

Of course, our analysis is subject to several limitations and boundary conditions, discussion of which highlights opportunities for additional research. For example, reinforcing spillover effects might be especially likely in our empirical context, since LEED was rapidly emerging as the *de facto* standard for green building certification and many private developers could reasonably expect that green building would yield direct economic benefits in the form of energy savings and increased demand. Moreover, governments are especially large customers in the construction services sector. Further research is needed to examine the extent to which public procurement rules influence private purchasing in mature markets where governments account for a smaller share of total demand.

Another caveat is that we do not measure the environmental performance implications of increased LEED adoption. While engineering studies suggest that LEED certification is correlated with increased energy efficiency, future research might examine the environmental impact of public green building policies.

Finally, since our findings suggest that government procurement policies can catalyze the adoption of a privately developed certification scheme, one might ask whether governments typically choose the “right” standard? In the case of LEED, it is not clear whether (a) municipal green building policies promoted lock-in to a particular standard (the leading alternative was the EPA’s Energy Star label) or (b) increasing returns simply led private and public actors to coalesce around the most popular measurement system at the time. Nevertheless, our LEED accreditation results show that government purchasing policies can promote standard-specific investments by various third parties, such as architects, contractors, and suppliers of green building materials. This both points to procurement policies as an effective policy tool and highlights the potential dangers of lock-in to a government-selected standard, particularly if it was developed by firms hoping to preempt more stringent regulation. The question of how government should be involved in the *ex ante* development of voluntary standards that might later provide the basis for procurement policies is another intriguing topic for future research.

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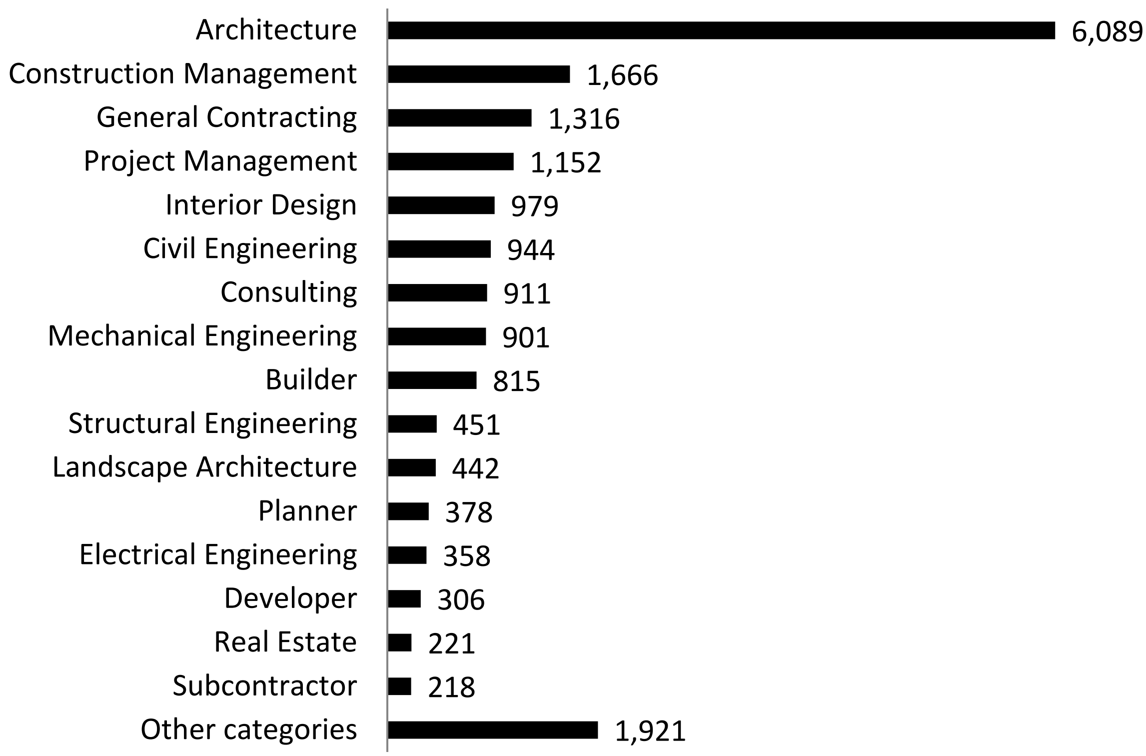
Figure 1. LEED Accredited Professionals by Occupation in California in 2010



Note: Based on data provided by the Green Building Certification Institute.

Table 1. Summary Statistics

Variable name	Definition	Mean	SD	Min	Max
Panel A: City-level Variables					
Total Private LEED Registrations	Total private LEED-registered buildings during 2001-2008	1.64	5.57	0.0	87.0
Total Private LEED Registrations per Capita	Total private LEED-registered buildings during 2001-2008 / City population in 10,000s	0.84	4.72	0.0	82.0
Total LEED Accredited Professionals	Total LEED Accredited Professionals during 2001-2008	7.51	27.38	0.0	416.0
Total LEED Accredited Professionals per Capita	Total LEED Accredited Professionals during 2001-2008 / City population in 10,000s	2.38	8.50	0.0	163.9
Total Government LEED Registrations	Total government LEED-registered buildings during 2001-2008	0.29	0.94	0.0	12.0
Total Government LEED Registrations per Capita	Total government LEED-registered buildings during 2001-2008 / City population in 10,000s	0.09	0.50	0.0	7.6
Green Policy Adopter	Focal city adopted green building policy by 2008 (dummy)	0.04	0.19	0.0	1.0
Green Policy Adopter Neighbor	Focal city borders a city that adopted green building policy by 2008 (dummy)	0.15	0.36	0.0	1.0
Exposure to Policy	Years elapsed by 2008 since focal city adopted green building policy	0.13	0.74	0.0	8.0
Exposure to Neighbor's Policy	Years elapsed by 2008 since neighbor city adopted green building policy	0.74	1.95	0.0	8.0
Prius Share	Toyota Prius as percent of all car registrations in 2008 (x 100)	0.54	0.59	0.0	3.7
Green Ballot Share	Percent of votes in favor of green ballot measures	60.73	14.76	19.9	100.0
LCV Senate Score	League of Conservation Voters score in 2001 of the city's State Senate representatives	39.31	42.16	0.0	100.0
LCV Assembly Score	League of Conservation Voters score in 2001 of the city's State Assembly representatives	40.55	42.02	0.0	100.0
Population	City population in 2000 (10,000s)	2.99	5.10	0.0	46.2
Total New Buildings	Total nonresidential construction starts during 2003-2007	26.21	54.71	0.0	869.0
Total New Buildings per Capita	Total nonresidential construction starts during 2003-2007 / City population in 10,000s	12.06	18.42	0.0	204.0
College	Percent college-educated in 2000	23.01	16.59	0.6	89.4
Income	Median household income in 2000 ($10,000s)	4.80	2.17	0.0	20.0
Log(Employment)	Employment in all SICs in 2001	8.07	2.06	0.0	13.9
Log(FIRE Employment)	Employment in Finance, Insurance & Real Estate SICs in 2001	4.89	2.50	0.0	11.41
Alternative Fuel Stations	Number of alternative-fuel filling stations in 2003	0.17	0.66	0.0	8.0
Solar Projects	Number of residential, commercial, and government solar installations by 2006	0.12	0.53	0.0	8.0
Panel B: City-year Variables					
Annual Private LEED Registrations	New private LEED-registered buildings this year	0.20	1.32	0.0	52.0
Annual Government LEED Registrations	New government LEED-registered buildings this year	0.04	0.24	0.0	6.0
Annual LEED Accredited Professionals	New LEED Accredited Professionals this year	0.94	5.21	0.0	160.0
City Adopted Green Policy	Focal city adopted policy by this year	0.02	0.13	0.0	1.0
Neighbor Adopted Green Policy	Neighbor city adopted policy by this year	0.09	0.29	0.0	1.0
Years Since City Adopted Green Policy	Years since focal city adopted procurement policy	0.03	0.27	0.0	4.0
Years Since Neighbor Adopted Green Policy	Years since neighbor city adopted procurement policy	0.25	0.90	0.0	4.0
Annual New Buildings	Nonresidential construction starts this year	5.90	12.90	0.0	208.7
Annual New Buildings per Capita	Nonresidential construction starts this year / City population in 10,000s	2.74	7.04	0.0	200.0

Notes: Panel A provides summary statistics for a cross-section of 735 California cities. Panel B reports annual variables for 5,880 city-year observations. Both panels exclude Los Angeles, San Diego, San Jose, and San Francisco.

Table 2. Covariate Balance in Full and Matched Samples

	Panel A			Panel B			Panel C		
Sample	Full sample			Green policy adopter cities and matched controls			Green policy adopter neighboring cities and matched controls		
Weighting	No weights			Weighted			Weighted		
	Green policy adopters	All non-adopters	t-stat	Green policy adopters	Matched controls	t-stat	Green policy adopter neighbors	Matched controls	t-stat
Prius Share	0.93	0.53	**3.62**	0.86	0.80	0.41	0.71	0.72	0.05
Green Ballot Share	72.26	60.25	**4.35**	71.08	68.27	1.09	68.08	65.98	1.36
LCV Senate Score	63.03	38.29	**3.12**	63.00	51.64	0.95	61.00	62.33	0.21
LCV Assembly Score	64.97	39.32	**3.25**	60.92	58.12	0.26	57.11	56.85	0.04
Alternative Fuel Stations	0.90	0.14	**6.23**	0.96	0.85	0.27	0.10	0.15	0.96
Solar Projects	0.86	0.09	**7.96**	0.96	0.74	0.71	0.07	0.08	0.02
Population	14.36	2.53	**13.68**	13.70	13.51	0.06	3.86	3.71	0.33
Total New Buildings	140.79	21.59	**12.64**	139.77	109.53	0.78	25.94	31.26	1.01
Total New Buildings per Capita	10.62	12.20	0.45	10.83	9.98	0.49	10.22	9.73	0.25
College	35.24	22.50	**4.09**	34.53	34.17	0.10	31.06	29.85	0.39
Income	5.58	4.77	**1.97**	5.70	5.83	0.33	5.98	6.04	0.12
Log(Employment)	10.97	7.96	**8.05**	10.92	10.68	0.94	9.25	9.00	1.29
Log(FIRE Employment)	8.22	4.76	**7.57**	8.19	8.15	0.11	6.44	6.13	1.22
Cities	29	697		26	180		80	291	

Notes: Unit of analysis is a city. Panel A reports means of each variable and t-statistic from unweighted OLS regression of the variable on *Green Policy Adopter* dummy. Panels B and C report CEM-weighted means of each variable and the t-statistic from CEM-weighted OLS regression of the variable on *Green Policy Adopter* dummy (middle panel) or *Green Policy Adopter Neighbor* dummy (right panel). CEM weights are described in Iacus, King, and Porro (2012) and discussed in the text. T-statistics corresponding to p < 10% are in boldface.

Table 3. Effects of Green Building Procurement Policies on LEED Registrations and Accredited Professionals: Cross-sectional Regression Results

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
Outcome	Total Private LEED Registrations		Total LEED Accredited Professionals	
	(1)	(2)	(3)	(4)
Exposure to Policy	2.11 [0.69]***		9.17 [4.75]*	
Exposure to Neighbor's Policy		0.15 [0.07]**		0.71 [0.29]**
Observations (cities)	206	371	206	371
CEM-weighted mean outcome	7.36	1.28	40.81	7.20
R-squared	0.56	0.28	0.39	0.41

Notes: CEM-weighted OLS regression coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city; see Table 2 for the number of treated and control units in the matched samples. All models also control for *Prius Share*, *Total New Buildings*, *Population*, *College*, *Income*, *Green Ballot Share*, *LCV Senate Score*, and *LCV Assembly Score*.

Table 4. Effects of Green Building Procurement Policies on LEED Registrations and Accredited Professionals: City Fixed-effects Regression Results

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
Outcome	Annual Private LEED Registrations		Annual LEED Accredited Professionals	
	(1)	(2)	(3)	(4)
Years Since City Adopted Green Policy	1.13 [0.73]		3.63 [2.29]	
Years Since Neighbor Adopted Green Policy		0.05 [0.02]**		0.24 [0.08]***
Weighted avg. outcome	0.92	0.16	5.08	0.90
Observations (city-years)	1672	2968	1672	2968
Cities	209	371	209	371
R-squared	0.31	0.16	0.48	0.31
F-test for pre-policy trend differences				
F-statistic	0.30	1.72	0.53	2.88
P value	0.87	0.16	0.71	0.03

Notes: CEM-weighted OLS coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city-year. All models also include city fixed effects and year fixed effects and control for *Annual New Buildings*.

Table 5. Effects of Green Building Procurement Policy Interacted with Green Demographics on LEED Registrations

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
Outcome	Total Private LEED Registrations					
	(1)	(2)	(3)	(4)	(5)	(6)
Exposure to Policy	2.11 [1.05]*		1.91 [0.90]**		2.12 [0.91]**	
Exposure to Neighbor's Policy		0.14 [0.06]**		0.14 [0.06]**		0.13 [0.05]**
Exposure to Policy × Prius Share	-0.50 [0.69]					
Exposure to Neighbor's Policy × Prius Share		0.22 [0.15]				
Exposure to Policy × College			0.03 [0.05]			
Exposure to Neighbor's Policy × College				0.00 [0.00]		
Exposure to Policy × Green Ballot Share					0.02 [0.06]	
Exposure to Neighbor's Policy × Green Ballot Share						0.00 [0.00]
Prius Share	6.83 [1.26]***	0.59 [0.17]***				
College			0.25 [0.03]***	0.02 [0.01]***		
Green Ballot Share					0.20 [0.15]	0.02 [0.02]
Total New Buildings	0.06 [0.02]***	0.03 [0.01]***	0.06 [0.02]***	0.03 [0.01]***	0.04 [0.02]*	0.03 [0.01]***
Observations (cities)	209	371	209	371	206	371
R-squared	0.42	0.28	0.45	0.26	0.36	0.22

Notes: CEM-weighted OLS regression coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city. *Prius Share, College*, and *Green Ballot Share* are demeaned when included in interaction terms.

Table 6. OLS and IV Estimates of Indirect Network Effects

Sample	Cities without a green building procurement policy and with population > 20,000			
Outcome	Total Private LEED Registrations		Total LEED Accredited Professionals	
Estimation	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
log(Total of LEED Accredited Professionals within 25 miles)	0.29 [0.12]**	0.56 [0.20]***		
Total Private LEED Registrations			2.15 [0.19]***	3.06 [1.37]***
	First-stage coefficients and statistics			
log(Number of cities with green policy within 25 to 50 miles in 2008)		1.21 [0.18]***		
Total New Buildings				0.04 [0.01]***
F-test of excluded IVs		44.53***		17.72***
Observations (cities)	244	244	244	244
Endogeneity test (p-value)		0.08		0.49
R-squared	0.43	0.42	0.50	0.46

Notes: OLS and instrumental variable regressions, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city. All models include controls for *Prius Share*, *Green Ballot Share*, *Population*, *College*, and *Income*. Models 1 and 2 also control for *Total New Buildings.* Null hypothesis for endogeneity test is exogeneity of endogenous regressor.

APPENDIX

Table A-1. California Cities with a Green Building Policy by 2008

	City	In matched sample	Population (10,000s)
1	Los Angeles	No	369.49
2	San Diego	No	122.34
3	San Jose	No	89.50
4	San Francisco	No	77.67
5	Long Beach	Yes	46.15
6	Sacramento	Yes	40.70
7	Oakland	No	39.95
8	Anaheim	Yes	32.80
9	Stockton	Yes	24.38
10	Fremont	Yes	20.34
11	Glendale	Yes	19.50
12	Santa Clarita	Yes	15.07
13	Santa Rosa	Yes	14.76
14	Irvine	Yes	14.31
15	Sunnyvale	Yes	13.18
16	Corona	Yes	12.50
17	Costa Mesa	Yes	10.87
18	Berkeley	No	10.27
19	Santa Clara	Yes	10.24
20	Ventura	No	10.09
21	Richmond	Yes	9.92
22	Santa Barbara	Yes	9.23
23	Santa Monica	Yes	8.41
24	San Leandro	Yes	7.95
25	Carlsbad	Yes	7.82
26	Livermore	Yes	7.33
27	Alameda	Yes	7.23
28	Temecula	Yes	5.77
29	La Mesa	Yes	5.47
30	Cupertino	Yes	5.05
31	West Hollywood	Yes	3.57
32	Dublin	Yes	3.00
33	Cotati	Yes	0.65

Table A-2. Effects of Green Building Procurement Policies on Government LEED Registrations: Cross-sectional Regression Results

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
Outcome	Total Government LEED Registrations	
	(1)	(2)
Exposure to Policy	0.56	
	[0.22]**	
Exposure to Neighbor's Policy		0.06
		[0.02]**
Observations (cities)	206	371
CEM-weighted mean outcome	0.93	0.26
R-squared	0.41	0.16

Notes: CEM-weighted OLS regression coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city; see Table 2 for the number of treated and control units in the matched samples. All models also control for *Prius Share*, *Total New Buildings*, *Population*, *College*, *Income*, *Green Ballot Share*, *LCV Senate Score*, and *LCV Assembly Score*.

Table A-3. Effects of Green Building Procurement Policies on LEED Registrations and Accredited Professionals: Cross-sectional Robustness Test Regression Results

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
	Panel A: Unweighted OLS Regressions			
Outcome:	Total Private LEED Registrations		Total LEED Accredited Professionals	
	(A1)	(A2)	(A3)	(A4)
Exposure to Policy	1.97		9.97	
	[0.67]***		[3.99]**	
Exposure to Neighbor's Policy		0.16		0.66
		[0.07]**		[0.27]**
Observations (cities)	206	371	206	371
Unweighted mean outcome	4.05	0.83	17.95	4.48
R-squared	0.64	0.34	0.63	0.38
	Panel B: Weighted Poisson Regressions			
Outcome:	Total Private LEED Registrations		Total LEED Accredited Professionals	
	(B1)	(B2)	(B3)	(B4)
Exposure to Policy	0.16		0.14	
	[0.04]***		[0.04]***	
Exposure to Neighbor's Policy		0.08		0.09
		[0.04]**		[0.03]***
Observations (cities)	206	371	206	371
CEM-weighted mean outcome	7.36	1.28	40.81	7.20

Notes: Figures are regression coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. All models include controls for *Prius Share*, *Total New Buildings*, *College*, *Income*, *Green Ballot Share*, *LCV Senate Score*, and *LCV Assembly Score*; for Panel B, these are included in log form (log after adding 1). Unit of analysis is a city.

Table A-4. Effects of Green Building Procurement Policies on LEED Registrations and Accredited Professionals: Robustness Tests

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
	Panel A: Unweighted OLS Regressions			
Outcome:	Annual Private LEED Registrations		Annual LEED Accredited Professionals	
	(A1)	(A2)	(A3)	(A4)
Years Since City Adopted Green Policy	1.50		5.77	
	[0.69]**		[2.41]**	
Years Since Neighbor Adopted Green Policy		0.09		0.42
		[0.02]***		[0.11]***
Observations (city-years)	1672	2968	1672	2968
Cities	209	371	209	371
Mean outcome	0.50	0.10	2.22	0.56
R-squared	0.24	0.13	0.32	0.20
	Panel B: Weighted OLS Regressions			
Outcome:	Log Annual Private LEED Registrations		Log Annual LEED Accredited Professionals	
	(B1)	(B2)	(B3)	(B4)
Years Since City Adopted Green Policy	0.12		0.06	
	[0.06]*		[0.05]	
Years Since Neighbor Adopted Green Policy		0.02		0.03
		[0.01]**		[0.02]*
Observations (city-years)	1672	2968	1672	2968
Cities	209	371	209	371
CEM-weighted mean outcome	0.33	0.09	1.01	0.34
R-squared	0.55	0.22	0.70	0.44
	Panel C: Weighted OLS First-differenced Regressions			
Outcome:	Annual Private LEED Registrations		Annual LEED Accredited Professionals	
	(C1)	(C2)	(C3)	(C4)
City Adopted Green Policy	0.78		3.29	
	[0.76]		[1.92]*	
Neighbor Adopted Green Policy		0.06		0.25
		[0.02]**		[0.11]**
Observations (city-years)	1463	2597	1463	2597
Cities	209	371	209	371
CEM-weighted mean outcome	0.50	0.09	2.91	0.55
R-squared	0.12	0.04	0.35	0.17

Notes: OLS coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city-year. All models control for *Annual New Buildings* and include year fixed effects. All models in Panel A include city fixed effects but do not include CEM weights. All models in Panel B are estimated on log dependent variables and include city fixed effects and CEM weights. All models in Panel C are estimated using first-differences of the dependent variable and *Annual New Buildings* and include CEM weights.

Table A-5. Effects of Green Building Procurement Policy Interacted with Green Demographics on LEED Registrations: Robustness Tests

Sample	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls	Green policy adopter cities and matched controls	Green policy adopter neighboring cities and matched controls
Outcome	Total Private LEED Registrations					
	(1)	(2)	(3)	(4)	(5)	(6)
Exposure to Policy	4.18 [1.54]**		3.42 [1.45]**		2.72 [1.28]**	
Exposure to Neighbor's Policy		0.08 [0.02]***		0.06 [0.03]**		0.09 [0.07]
Exposure to Policy × Prius Share above sample median	-2.71 [1.91]					
Exposure to Neighbor's Policy × Prius Share above sample median		0.13 [0.11]				
Exposure to Policy × College above sample median			-1.84 [1.57]			
Exposure to Neighbor's Policy × College above sample median				0.14 [0.12]		
Exposure to Policy × Green Ballot Share above sample median					-0.69 [1.56]	
Exposure to Neighbor's Policy × Green Ballot Share above sample median						0.09 [0.15]
Prius Share above sample median	7.47 [1.66]***	0.84 [0.19]***				
College above sample median			7.15 [1.90]***	0.74 [0.16]***		
Green Ballot Share above sample median					3.34 [2.70]	0.49 [0.35]
Total New Buildings	0.06 [0.02]***	0.02 [0.01]***	0.06 [0.02]***	0.02 [0.01]***	0.05 [0.02]**	0.03 [0.01]***
Observations (cities)	209	371	209	371	209	371
R-squared	0.42	0.25	0.41	0.24	0.34	0.22

Notes: CEM-weighted OLS regression coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city.

Table A-6. Effects of Green Building Procurement Policy Interacted with City Size on LEED Accredited Professionals

Sample	Green policy adopter cities and matched controls			Green policy adopter neighboring cities and matched controls		
Outcome	Total LEED Accredited Professionals					
	(1)	(2)	(3)	(4)	(5)	(6)
Exposure to Policy	3.62	11.36	-3.48			
	[4.02]	[8.37]	[6.29]			
Exposure to Policy × log(City Population)	8.66					
	[5.72]					
Exposure to Policy × log(County Population)		0.75				
		[3.98]				
Exposure to Policy × log(Total New Buildings)			13.86			
			[6.60]**			
Exposure to Neighbor's Policy				0.85	0.96	0.31
				[0.37]**	[0.66]	[0.18]*
Exposure to Neighbor's Policy × log(City Population)				-0.24		
				[0.23]		
Exposure to Neighbor's Policy × log(County Population)					-0.14	
					[0.28]	
Exposure to Neighbor's Policy × log(Total New Buildings)						0.43
						[0.33]
log(City Population)	7.05			2.51		
	[2.21]***			[0.94]**		
log(County Population)		3.34			-0.35	
		[3.76]			[0.92]	
log(Total New Buildings)			10.93			3.92
			[4.30]**			[1.27]***
Observations (cities)	209	209	209	371	371	371
R-squared	0.20	0.10	0.36	0.07	0.02	0.18

Notes: CEM-weighted OLS coefficients, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city. *Population* and *Total New Buildings* are demeaned when included in interaction terms.

Table A-7. OLS and IV Estimates of Indirect Network Effects (All Cities)

Sample	All cities without green procurement policy			
Outcome	Total Private LEED Registrations		Total LEED Accredited Professionals	
Estimation	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)
log(Total of LEED Accredited Professionals within 25 miles)	0. 10 [0.05]*	0. 12 [0.07]*		
Total Private LEED Registrations			2.31 [0.18]***	3.06 [1.08]***
	First-stage coefficients and statistics			
log(Number of cities with green policy within 25 to 50 miles in 2008)		1.29 [0.13]***		
Total New Buildings				0.04 [0.01]***
F-test of excluded IVs		97.17***		17.70***
Observations (cities)	697	697	697	697
Endogeneity test (p-value)		0.66		0.46
R-squared	0.38	0.38	0.54	0.52

Notes: OLS and instrumental variable regressions, with robust standard errors clustered by county in brackets; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city. All models include controls for *Prius Share*, *Green Ballot Share*, *Population*, *College*, and *Income*. Models 1 and 2 also control for *Total New Buildings.* Null hypothesis for endogeneity test is exogeneity of endogenous regressor.